Inflation in developing countries: does central bank independence matter? New evidence based on a new data set

Jan-Egbert Sturm and Jakob de Haan*

Department of Economics, University of Groningen, The Netherlands

Version, March 2001

Abstract

We analyse whether central bank independence (CBI) affects inflation in developing countries. For this purpose we have constructed a new data set for the turnover rate (TOR) of central bank governors for a very large sample of countries, which also covers the 1990s. We find that once various control variables are included, the CBI proxy is often not significant. We also conclude that in those regressions in which the CBI proxy is significant, the coefficient of the TOR becomes significant only after high inflation countries are added to the sample.

JEL code: E58, E52
Key words: inflation, central bank independence

* Corresponding author: Department of Economics, University of Groningen, PO Box 800, 9700 AV Groningen, The Netherlands. Tel. 31-50-3633706. Fax 31-50-3633720. Email: j.de.haan@eco.rug.nl.

1. Introduction

Nowadays it is widely believed that a high level of central bank independence (CBI) coupled with some explicit mandate for the bank to restrain inflation are important institutional devices to assure price stability. The evidence in support of this view generally consists of cross-country regressions using proxies for CBI which are generally based on the statutes of the central bank (see Eijffinger and De Haan, 1996 and Berger et al., 2001 for surveys). Popular as the view referred to above may be, the empirical literature on the effects of central bank independence recently came under attack.

Firstly, the relevance of independence measures which are based on regulations on the position of the central bank is disputed if it comes to testing whether CBI is conducive to lower inflation (see e.g. Forder, 1996). Indeed, legal indices of central bank independence are often incomplete and noisy indicators of actual independence as laws cannot specify explicitly the limits of authority between central banks and the political authorities under all contingencies. And even when the laws are quite explicit, actual practice may deviate from them. Cukierman (1992) argues that legal independence measures may be a better proxy for actual independence in industrial countries than in developing countries. As an alternative, Cukierman (1992) and Cukierman et al. (1992) therefore developed a yardstick for central bank autonomy, which is not based on central bank laws but on the actual average term of office of the central bank governor. This indicator is based on the presumption that, at least above some threshold, a higher turnover of central bank governors indicates a lower level of independence. So far, quite a few studies have used this turnover rate of central bank governors (TOR) as indicator for CBI and conclude that there is a clear relationship between the TOR and the inflation performance of developing countries. One drawback of this literature is that almost all studies are based on the data provided by Cukierman (1992) and Cukierman et al. (1992) as this was the only data set available.[1]

Secondly, various authors have questioned whether CBI really matters, once other variables that may influence inflation are taken into account. For instance, using data for developing countries Campillo and Miron (1997) conclude that CBI plays no role in determining inflation outcomes, once other factors are held constant. They find that instead openness, political instability and proxies for government policy distortions are robustly related to inflation.[2] This conclusion may be criticised, as Campillo and Miron (1997) employ a legal indicator for CBI, which most previous studies have found to be unreliable for developing countries.

Thirdly, a few studies have sounded a warning that conclusions on the relationship between CBI and inflation are highly sensitive to influential observations. For instance, Temple (1998) finds that if high inflation countries are added to his sample of OECD and developing countries, the effect of CBI (proxied by Cukierman's (1992) legal index) on inflation disappears.

We employ a new data set for the turnover rate of central bank governors, which covers almost twice as many countries as the Cukierman data set and - in contrast to Cukierman's data set - also covers the 1990s. We use bivariate and multivariate cross-country models for inflation. We find that once various control variables are included, the CBI proxy is often not significant. Following the approach suggested by Temple (1998), we also conclude that in those regressions in which the CBI proxy is significant, the coefficient of the TOR becomes significant only after high inflation countries are added to the sample.

The remainder of the paper is organised as follows. Section 2 presents the model and our data. Section 3 offers our results. The final section contains some concluding comments.

2. The model and the data

Following Cukierman et al. (1992) we have used the transformed inflation rate D as dependent variable in order to reduce heteroscedasticity of the error in the regressions. D is defined as the inflation rate (p) divided by one plus the inflation rate:

$$D = p/(1 + p) \qquad (1)$$

So the transformed inflation rate (if positive) takes a value from 0 to 1. When inflation is 100% (p=1), D is 0.5. D has been calculated for each year, and subsequently the averages for the various estimation periods have been calculated as the simple means of these annual observations. The inflation rate is taken from the World Bank's *2000 World Development Indicators* CD-rom.

One serious drawback of many studies on the relationship between inflation and CBI is that control variables are often lacking. Following Campillo and Miron (1997), we therefore also estimate multivariate models. We include apart from indicators for CBI and openness (export and import as share of GDP), also political instability (proxied by the number of government transfers), the log of GDP per capita, a dummy for the exchange rate regime (one in case of a more or less fixed exchange rate) and the debt-to-GDP ratio as control variables.

Our indicator for openness (OPEN, defined as sum of export and import in relation to GDP) and the log of the level of GDP per capita (GDPCAP) are from the World Bank's *2000 World Development Indicators* CD-rom. Our proxy for political instability (PI, defined as the total number of government transfers) is from the new World Bank data set of political indicators.[3] An exchange rate dummy - which is one if the country had at the beginning and the end of the period a more or less fixed exchange rate regime - is used to examine the impact of the exchange rate regime. This variable is denoted as XRATE. It is constructed using information reported in the IMF's Annual Report. The external debt-to-GDP ratio at the beginning of the estimation period (DEBT) is also from the World Bank. So the estimated cross-section model is:[4]

$$D = c_0 + c_1 CBI + c_2 OPEN + c_3 GDPCAP + c_4 PI + c_5 XRATE + c_6 DEBT \qquad (2)$$

We use indicators for central bank independence based on the turnover rate of central bank governors. Apart from the TOR as provided by Cukierman et al. (1992), we employ a new data set.[5] Based on information gathered from central banks and the IMF's *International Financial Statistics* (IFS) we have constructed the turnover rate of the central bank governors in more than 80 developing countries for various sample periods. To enable comparison with the data of Cukierman (1992), we calculated the turnover rates for two periods: 1980-1989 and 1990-1998.[6] For the first period Cukierman (1992) also provides turnover rates, albeit for a substantially smaller sample of countries. Detailed information is shown in Table A1 in the Appendix.

3. Central bank independence and inflation in developing countries

We focus on the relationship between CBI and inflation in developing countries for two periods: 1980-1989 and 1990-1998. The first period is chosen on the basis of the availability of the data of Cukierman (1992). We not only estimate simple bivariate regressions, but following Campillo and Miron (1997) also estimate models that include various control variables (see Section 2). Unfortunately, this reduces the number of observations.

We start with the simple bivariate model using our new data set for the TOR. Row 1 of table 1 shows the OLS estimation results for the period 1980-89. The coefficient of the TOR is highly significant. The same results show up when we employ Cukierman's TOR (row 2 of table 1). Also for the period 1990-98 the coefficient of our CBI proxy differs significantly from zero. These results are very much in line with almost all previous studies, which found that there is a clearly significant positive relationship between the TOR and inflation in developing countries.

[Insert Table 1 about here]

We checked whether the relationship between inflation and CBI is linear using the RESET (Regression Error Specification Test). This test means regressing the in-sample residuals on the same regressors as in the original linear regression and on powers of the in-sample forecasts and testing for the joint significance of the coefficients performing an F-test (see e.g. Granger and Terasvita, 1993).[7] These tests clearly lead to the conclusion that the relationship is linear.[8]

Next, we checked for the influence of high inflation countries. A very simple way to visualise the role of these countries - as suggested by Temple (1998) - is to order the countries according to their level of inflation and then add countries one by one to the sample. Figure 1 shows the estimated coefficients and the band of plus and minus two times the standard error for the regressions reported in rows 1-3 of table 1. The figures clearly show that only after the high inflation countries are included in the sample, the coefficient of the TOR index becomes significant. In other words, the conclusion that CBI matters for inflation is driven by a limited number of observations.

[Insert Figure 1 about here]

Next, we turn to the multivariate models. The lower part of table 1 presents the regressions. As we are mainly interested in the effect of CBI on inflation, we only report the estimated coefficients for the TOR.[9] It follows that in the regression for the period 1980-89 the coefficient of our proxy for CBI is not significantly different from zero (row 4 of table 1). When we use Cukierman's TOR instead, its coefficient is significantly different from zero, although only at the 10 per cent level (row 5). For the 1990s we do not find a significant effect of our TOR on inflation for the multivariate model (row 6). The variables that turn out to be significant in all regressions are openness, the exchange rate regime (except in the regression with Cukierman's TOR) and the debt ratio (see table A2 in the Appendix).

Figure 2 shows the estimated coefficients and the band of plus and minus two times the standard error for the regressions reported in rows 4-6 of table 1. Again, it clearly follows that Cukierman's TOR only becomes significant once high inflation countries are added to the sample. Although the coefficients of our proxy for CBI never become significant, the same important role of high inflation observations can be discerned.

[Insert Figure 2 about here]

4. Conclusions

In this paper we have re-examined the relationship between central bank independence and inflation in developing countries. We extend the existing literature in three ways. Firstly, we present a new data set for the turnover rate of central bank governors in a very large sample of countries, which also covers the 1990s. Secondly, we employ turnover rates in a multivariate model. Previous studies which used control variables (notably Campillo and Miron, 1997 and Temple, 1998) employed legal indicators for CBI, which are widely believed to be less relevant for developing countries. Third, we examined the role of high inflation countries, following the approach suggested by Temple (1998). Our results are in sharp contrast to most previous studies as we find that once various control variables are included, the CBI proxy is often not significant. We also conclude that in those regressions in which the CBI proxy is significant, the coefficient of the TOR becomes significant only after high inflation countries are added to the sample.

References

Beck, T., G. Clarke, A. Groff, P. Keefer and P. Walsh, 1999, New tools and new tests in comparative political economy: The Database of Political Institutions, July 1999.

Berger, H., S.C.W. Eijffinger, and J. de Haan, 2001, 'Central Bank Independence: An Update of Theory and Evidence', Journal of Economic Surveys, Vol. 15, 3-40.

Campillo, M. and Miron, J.A., 1997, 'Why does inflation differ across countries?' In: Romer, C.D., Romer, D.H. (Eds.), Reducing Inflation: Motivation and Strategy, Chicago: University of Chicago Press.

Cukierman, A., 1992, Central bank strategy, credibility, and independence, Cambridge: MIT Press,.

Cukierman, A., S.B. Webb and B. Neyapti, 1992, 'Measuring the independence of central banks and its effects on policy outcomes', The World Bank Economic Review, Vol. 6, 353-398.

Eijffinger, S.C.W. and J. de Haan, 1996, The political economy of central-bank independence, Princeton Special Papers in International Economics, No. 19.

Forder, J., 1996, 'On the assessment and implementation of `institutional' remedies', Oxford Economic Papers, Vol. 48, 39-51.

Granger, C.W.J, and T. Terasvirta, 1993, Modelling Non-linear Economic Relationships, Oxford: Oxford University Press.

Haan, J. de and Kooi, W., 2000, 'Does central bank independence really matter? New evidence for developing countries using a new indicator', Journal of Banking and Finance, Vol. 24, 643-664.

Lee, T.H., H. White, and C.W.J. Granger. 1993, 'Testing for Neglected Non-linearities in Time Series Models', Journal of Econometrics, Vol. 56, 269-290.

Romer, D., 1993, 'Openness and inflation: Theory and evidence', Quarterly Journal of Economics, Vol. 108, 869-904.

Temple, J., 1998, 'Central bank independence and inflation: good news and bad news', Economics Letters, Vol. 61, 215-219.

White, H., 1980, 'A heteroscedasticity-consistent covariance matrix estimator and a direct test for heteroscedasticity', Econometrica, Vol. 48, 817-838.

Table 1. OLS regressions for the relationship between CBI and inflation

Period and CBI proxy:	CBI coefficient	t-statistic	R2 (adj.)	No. obs
Bivariate models				
(1) 1980-89; new TOR	30.55	2.57 ***	0.20	76
(2) 1980-89; Cukierman TOR	41.26	3.01 **	0.22	45
(3) 1990-98; new TOR	23.26	2.30 **	0.11	76
Multivariate models				
(4) 1980-89; new TOR	14.61	1.33	0.46	55
(5) 1980-89; Cukierman TOR	24.04	1.70 *	0.53	29
(6) 1990-98; new TOR	15.34	1.52	0.27	63

Note: heteroscedasticity-consistent t-ratios (White, 1980). In the regressions for the multivariate model political instability, openness, the log of GDP per capita, a dummy for the exchange rate regime and the debt-to-GDP ratio are included as control variables. *,**,*** denote 10%, 5%, and 1% significance levels, respectively.

Figure 1: Rolling regressions for the bivariate relationship between CBI and inflation

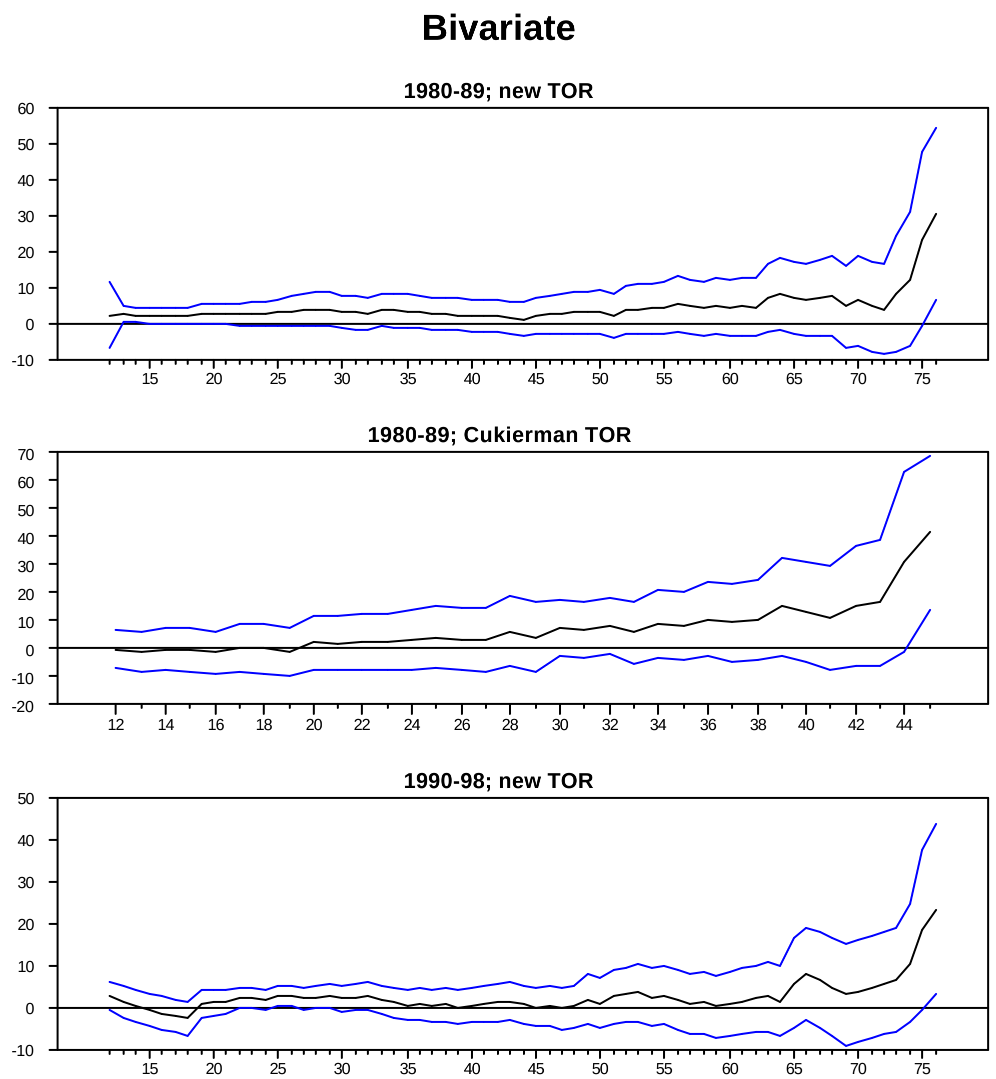
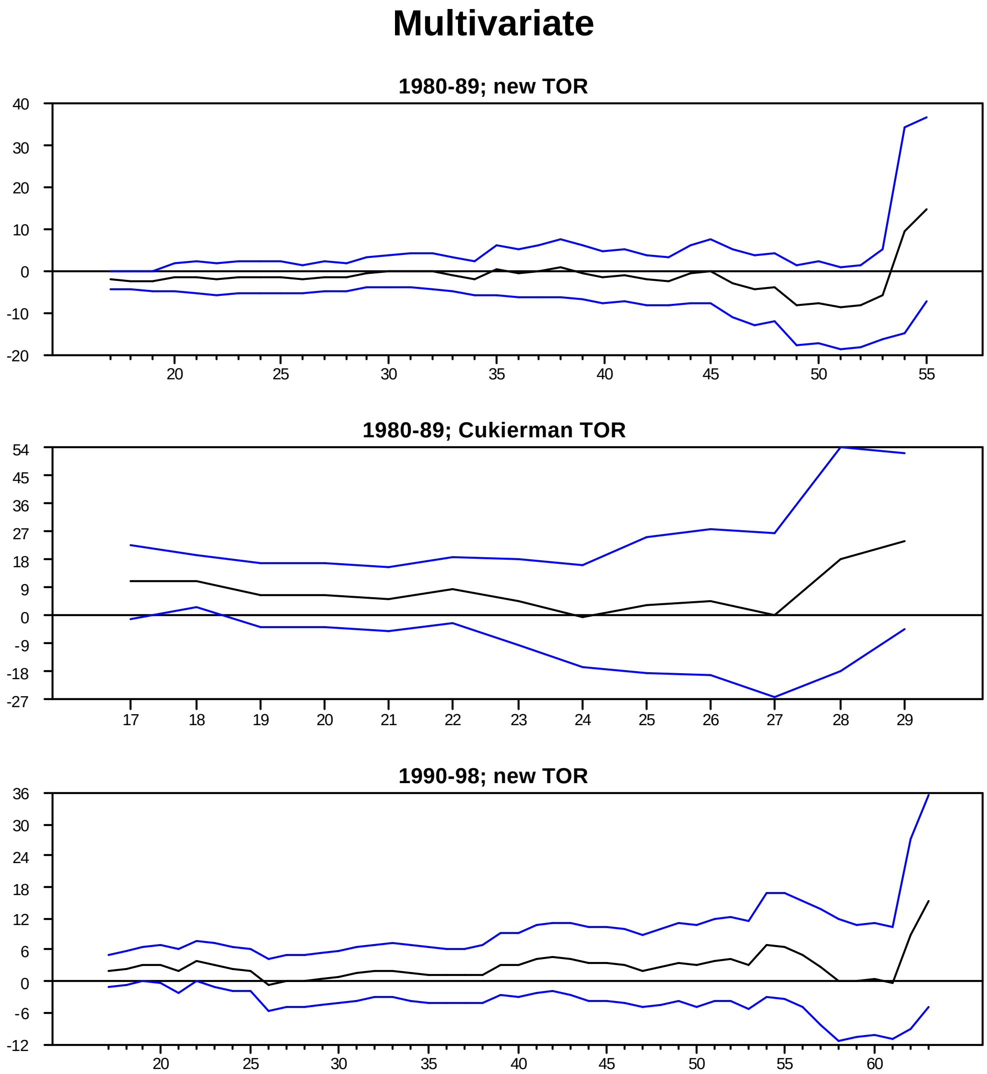


Figure 2: Rolling regressions for the multivariate relationship between CBI and inflation



Appendix.

Table A1. Turnover rates of central bank governors, 1980-89 and 1990-98.

Country	Cukierman 1980-89	New 1980-89	New 1990-98
Algeria	NA	0.30	0.00
Argentina	1.00	1.00	0.25
Bahamas	0.20	0.10	0.25
Bahrain	NA	0.00	0.00
Bangladesh	NA	0.20	0.25
Barbados	0.10	0.20	0.25
Belize	NA	0.30	0.00
Botswana	0.40	0.10	0.25
Brazil	0.80	1.10	0.75
Burundi	NA	0.20	0.00
Cape Verde	NA	0.10	0.00
Chile	0.80	0.40	0.25
China	0.30	NA	NA
Colombia	0.20	0.20	0.00
Costa Rica	0.40	0.30	0.50
Cyprus	NA	0.00	0.00
Djibouti	NA	0.10	0.25
Dom Rep.	NA	0.60	0.00
Ecuador	NA	0.70	0.75
Egypt	0.30	0.30	0.00
El Salvador	NA	0.30	0.25
Ethiopia	0.10	0.20	0.50
Fiji	NA	0.10	0.00
Gambia	NA	0.30	0.00
Ghana	0.20	0.10	0.25
Greece	0.20	0.30	0.00
Guatemala	NA	0.80	0.25
Guinea	NA	0.10	0.25
Guinea Bissau	NA	0.10	0.50
Haiti	NA	1.00	0.25
Honduras	0.10	0.20	0.25
Hungary	0.10	NA	NA

Country	Cukierman 1980-89	New 1980-89	New 1990-98
India	0.30	0.40	0.25
Indonesia	0.20	0.20	0.25
Iran	NA	0.30	0.00
Israel	0.20	NA	NA
Jamaica	NA	0.50	0.25
Jordan	NA	0.20	0.25
Kenya	0.20	0.20	0.00
Korea South	0.50	0.40	0.50
Kuwait	NA	0.10	0.00
Lebanon	0.10	0.20	0.00
Lesotho	NA	0.20	0.25
Libya	NA	0.30	0.25
Madagascar	NA	0.20	0.25
Malawi	NA	0.30	0.25
Malaysia	NA	0.20	0.25
Maldives	0.20	0.00	0.00
Malta	0.20	0.20	0.25
Mauritius	NA	0.00	0.25
Mexico	0.30	0.00	0.25
Morocco	0.20	0.20	0.00
Mozambique	NA	0.20	0.00
Nepal	0.10	0.20	0.25
Nicaragua	0.40	0.50	0.25
Nigeria	0.10	0.10	0.00
Pakistan	0.30	0.50	0.00
Panama	0.20	NA	NA
Paraguay	NA	0.30	0.50
Peru	0.30	0.50	0.00
Philippines	0.20	0.20	0.00
Poland	0.50	NA	NA
Portugal	0.30	NA	NA
Qatar	0.00	0.10	0.00
Romania	0.20	NA	NA
Saudi Arabia	NA	0.00	0.00
Seychelles	NA	0.10	0.25

Country	Cukierman 1980-89	New 1980-89	New 1990-98
Singapore	0.60	0.30	0.25
Solomon Islands	NA	0.10	0.00
South Africa	0.20	0.10	0.00
Sri Lanka	NA	0.20	0.25
Sudan	NA	0.44	0.50
Surinam	NA	0.20	0.25
Swaziland	NA	0.10	0.25
Syria	NA	0.10	0.25
Taiwan	0.20	NA	NA
Tanzania	0.10	0.20	0.25
Thailand	0.10	0.10	0.75
Trin & Tobago	NA	0.20	0.25
Tunisia	NA	0.30	0.00
Turkey	0.40	0.30	0.25
Uganda	0.20	0.20	0.00
Uruguay	0.30	0.30	0.50
Vanuatu	NA	0.40	0.25
Venezuela	0.50	0.50	0.00
Western Samoa	0.56	0.20	0.00
Yugoslavia	0.20	NA	NA
Zaire	0.20	0.30	NA
Zambia	0.50	0.40	0.25
Zimbabwe	0.10	0.10	0.00

Table A2. Multivariate cross-country models of inflation in developing countries

Variable:	new TOR		Cukierman TOR		new TOR	
	1980-89		1980-89		1990-98	
	Coeff.	t-stat.	Coeff.	t-stat.	Coeff.	t-stat.
Constant	-14.30	(-1.39)	-15.15	(-1.02)	4.92	(0.56)
CBI	14.61	(1.33)	24.04	(1.70)	15.34	(1.52)
GDPCAP	3.82	(2.30)	5.40	(2.17)	1.88	(1.36)
OPEN	-0.13	(-2.56)	-0.40	(-3.91)	-0.09	(-2.77)
PI	16.15	(1.15)	6.04	(0.21)	-2.14	(-0.25)
XRATE	-5.17	(-2.00)	-1.08	(-0.23)	-7.15	(-3.24)
DEBT	0.16	(5.70)	0.16	(3.97)	0.04	(3.65)
No. obs.	55		29		63	
R2 (adj.)	0.46		0.53		0.27	

Notes

[1] Furthermore, the Cukierman data refer to a small sample of countries only.

[2] Their results on openness confirm earlier findings of Romer (1993), who argues that monetary surprises cause the real exchange rate to depreciate, and since real depreciations has the most harmful effects in more open economies, the benefits of unexpected inflation are a decreasing function of the degree of openness. In the absence of binding pre-commitments, monetary authorities in more open economies are therefore expected to expand less on average.

[3] See Beck et al. (1999). The variable we use is labeled STABNS2 that is defined as the percent of veto players who drop from the government in any given year. Veto players are the president, the largest government party, and the largest party in the Senate; for parliamentary systems, veto players are defined as the Prime Minister and the biggest three coalition members. If there is no legislature, an unelected legislature, only 1 candidate, or 1 party to choose from during elections this index is based only on changes in the chief executive.

[4] Given the limited variation over time in the TOR variable, a panel data approach is not helpful in the present context.

[5] Part of this data set (referring to 1980-89 only) has already been used in De Haan and Kooi (2000).

[6] The underlying data (on an annual basis) are available upon request.

[7] Due to the possibly large degree of multicollinearity in these powers, Lee, White and Granger (1993) suggest using reconstructed powers on principal components, excluding the first and largest one.

[8] The test statistic for row 1 in table is e.g. 1.947 (significance level: 0.15).

[9] See table A2 in the Appendix for further details.